

# Mean Reversion of Short-run Interest Rates in Emerging Countries\*

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## Abstract

In this paper we examine the stochastic behavior of short-run interest rates in several emerging countries using fractional integration techniques. We allow for a much richer flexibility in the dynamic behavior of the series than the classical representations based on  $I(0)$  or  $I(1)$  processes. It appears that for Singapore and Thailand nominal interest rates are mean-reverting, whereas for Mexico, Malaysia, the Philippines, and Korea, the presence of a unit-root test depends on the assumptions regarding the residuals' autocorrelation. The results also suggest that uncovered interest parity (UIP) can only hold for two emerging countries. For the other countries, the stabilization policies in the aftermath of the currency crises have led to the rejection of the UIP hypothesis.

## 1. Introduction

The analysis of interest rate persistence is a major question in the empirical literature. Rose (1988), Stock and Watson (1988), Campbell and Shiller (1991), and Wu and Chen (2001) have applied a number of unit-root tests to determine whether short-run interest rates are stationary or not. The mean-reverting property of the series has major consequences. First, in terms of modeling strategies, it turns out that vector-error correction (VEC) or vector autoregression (VAR) in differences are not necessary to model the dynamics of short-run interest rates. A simple VAR in levels appears sufficient to represent the dynamics of short-run interest rates. Secondly, the rejection of a unit-root test in short-run interest rates sheds some light on the empirical investigation of two major relationships in macroeconomics: the Fisher hypothesis (FH) and the uncovered interest parity hypothesis (UIP). If interest rates and inflation are found to be nonstationary (or  $I(1)$ ), a long-run version of the FH can be tested within a co-integration framework (Mishkin, 1992). Moreover, in such a case, inflation expectations have a permanent effect on interest rates, in contrast to *ex-ante* real interest rate shocks. In the case where interest rates are no longer infinitely persistent, the previous analysis no longer applies. Hence, the presence of a long-run FH is rejected, whereas a short-run FH can be accepted. The stationarity of nominal short-run interest rates also has consequences on the empirical validity of UIP. As nominal bilateral exchange rates are difference-stationary, the validity of the UIP relation requires mean-reverting nominal short-run interest rates.

Previous empirical studies conclude that short-run interest rates are mean-reverting in Europe and in the US (e.g. Rose, 1988; Stock and Watson, 1988; Wu and Chen, 2001),

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implying the rejection of a long-run FH and the possible investigation of the UIP relationship. Such a result has repercussions on the monetary policymaking. The validity of UIP would indicate that international flows of capital would constrain the central bank's ability to fix interest rates.

Nevertheless, in the case of emerging countries, interest rates are often used as an instrument for stabilization policy (for example, to adjust the nominal exchange rates) to prevent capital outflows during financial crises. Thus, interest rates can cease to be mean-reverting. In that case, the UIP condition cannot hold (especially if a country with fixed exchange rates is considered), whereas the empirical validity of the long-run FH via cointegration analysis is relevant.

In this paper we analyze the mean-reversion properties of short-term interest rates for a panel of emerging countries. In particular, we consider one Latin American country (Mexico), with periods of fixed exchange rates, and five South Asian countries (Thailand, Korea, Malaysia, the Philippines, and Singapore), with floating exchange rates. We use the fractional integration framework that also appears in Shea (1991), Backus and Zin (1993), Connolly and Guner (1999), and Duan and Jacobs (2001). This framework provides a more general and flexible alternative for investigating data dynamics than the traditional stationary (I(0)) and nonstationary (I(1)) approaches. As it allows one to consider non-integer differences to be applied in raw time series, the testing procedure proposed by Robinson (1994) for fractional integration will be used to determine the degree of integration of the series.

## 2. Fractional Integration and Mean Reversion

Modeling macroeconomic time series remains controversial. Initially, deterministic approaches based on linear (or quadratic) time functions were proposed but they were shown to be inappropriate in many cases, in particular if the trend changes over time. Next, and especially after the seminal paper of Nelson and Plosser (1982), stochastic approaches based on first (or second) differences of the data became popular. Nelson and Plosser, following the work and ideas of Box and Jenkins (1970), showed that many US macroeconomic variables could be specified in terms of unit roots. Using tests of Fuller (1976) and Dickey and Fuller (1979), they were unable to reject the unit-root hypothesis in practically all of the series that they analyzed. However, unit-root models can be viewed as a special case of a more general class of processes called long memory processes, due to their ability to display significant dependence between observations widely separated in time. A popular technique to analyze fractionally integrated models is through the fractional differencing operator  $(1 - L)^d$ , where:

$$(1 - L)^d = \sum_{k=0}^{\infty} (-1)^k \binom{d}{k} L^k = 1 - dL + \frac{d(d-1)}{2} L^2 - \frac{d(d-1)(d-2)}{6} L^3 + \dots$$

and  $L$  is the lag operator ( $Lx_t = x_{t-1}$ ). To illustrate this technique in case of a scalar time series  $x_t$ ,  $t = 1, 2, \dots$ , suppose that  $u_t$  is an unobservable covariance stationary sequence with spectral density that is bounded and larger than zero at any frequency, and

$$(1 - L)^d x_t = u_t, \quad t = 1, 2, \dots \quad (1)$$

The process  $u_t$  could itself be a stationary and invertible ARMA sequence, when its autocovariances decrease exponentially. However, their autocovariances could decrease much slower than exponentially. When in (1)  $d = 0$ ,  $x_t = u_t$  and thus  $x_t$  is

“weakly autocorrelated,” also called “weakly dependent.” If  $0 < d < 0.5$ ,  $x_t$  is still stationary but its lagged  $j$  autocovariance  $\gamma_j$  decreases very slowly, like the power law  $j^{2d-1}$  as  $j \rightarrow \infty$  and so the  $\gamma_j$  are nonsummable. We say then that  $x_t$  has long memory given that its spectral density  $f(\lambda)$  is unbounded at the origin. Finally, if in (1)  $d$  increases such that it lies in the half-open interval  $]0.5, 1[$ ,  $x_t$  can be viewed as becoming “more nonstationary” in the sense that the variance of the partial sums increases in magnitude. Because this also holds for  $d > 1$ , a large class of nonstationary processes may be described by (1) with  $d \geq 0.5$ . The distinction between  $I(d)$  with different values of  $d$  is also important from an economic point of view: if  $d < 1$ , the process is mean-reverting, with shocks affecting the system, but the variable  $x_t$  returns to its original level somewhere in the future. On the other hand,  $d \geq 1$  means that the series is nonstationary and not mean-reverting. Thus, the fractional differencing parameter  $d$  plays a crucial role in our understanding of the economy, and economic stabilization policy. In particular, if a variable exhibits a unit root, any shock to the economic system will have a permanent effect, such that a policy action will be required to bring the variable back to its original long-term target. On the other hand, if  $d$  is smaller than 1, fluctuations will be transitory and therefore there is less need for policy intervention, since the series will return to its trend anyway.

Long memory processes were initially introduced by Granger (1980, 1981), Granger and Joyeux (1980), and Hosking (1981), and were theoretically justified by Robinson (1978) and Granger (1980). Similarly, Ciozbek-Georges and Mandelbrot (1995), Taquq et al. (1997), Chambers (1998), and Lippi and Zaffaroni (1999) use aggregation to motivate long memory processes as well, while Parke (1999) uses a closely related discrete time-error duration model. The possibility of long memory in interest rates has been examined by Shea (1991), Backus and Zin (1993), Connolly and Guner (1999), Duan and Jacobs (2001), and others. However, a proper study of these series in terms of estimation and testing of  $I(d)$  models still needs to be pursued. In this article, we claim that short-run interest rates in some emerging countries may be described in terms of  $I(d)$  statistical processes.

### 3. A Testing Procedure for $I(d)$ Statistical Models

Robinson (1994) proposes a Lagrange multiplier (LM) test for the null hypothesis:

$$H_0: d = d_0 \quad (2)$$

in a model given by:

$$y_t = \beta' z_t + x_t, \quad t = 1, 2, \dots, \quad (3)$$

and (1), for any real value  $d_0$ , where  $y_t$  is the time series we observe,  $\beta = (\beta_1, \dots, \beta_k)$  is a  $(k \times 1)$  vector of unknown parameters, and  $z_t$  is a  $(k \times 1)$  vector of deterministic regressors that may include an intercept (e.g.  $z_t \equiv 1$ ), or an intercept and a linear time trend (in the case of  $z_t = (1, t)'$ ). Specifically, the test statistic is given by:

$$\hat{r} = \frac{T^{1/2}}{\hat{\sigma}^2} \hat{A}^{-1/2} \hat{a}, \quad (4)$$

where  $T$  is the sample size and

$$\hat{a} = \frac{-2\pi}{T} \sum_{j=1}^{T-1} \psi(\lambda_j) g(\lambda_j; \hat{\tau})^{-1} \mathbf{I}(\lambda_j); \quad \hat{\sigma}^2(\hat{\tau}) = \sigma^2(\hat{\tau}) = \frac{2\pi}{T} \sum_{j=1}^{T-1} g(\lambda_j; \hat{\tau})^{-1} \mathbf{I}(\lambda_j);$$

$$\hat{A} = \frac{2}{T} \left( \sum_{j=1}^{T-1} \psi(\lambda_j)^2 - \sum_{j=1}^{T-1} \psi(\lambda_j) \hat{\varepsilon}(\lambda_j)' \times \left( \sum_{j=1}^{T-1} \hat{\varepsilon}(\lambda_j) \hat{\varepsilon}(\lambda_j)' \right)^{-1} \times \sum_{j=1}^{T-1} \hat{\varepsilon}(\lambda_j) \psi(\lambda_j) \right);$$

$$\psi(\lambda_j) = \log \left| 2 \sin \frac{\lambda_j}{2} \right|; \quad \hat{\varepsilon}(\lambda_j) = \frac{\partial}{\partial \tau} \log g(\lambda_j; \hat{\tau}); \quad \lambda_j = \frac{2\pi j}{T}; \quad \hat{\tau} = \arg \min_{\tau} \sigma^2(\tau).$$

$I(\lambda_j)$  is the periodogram of  $u_t$  evaluated under the null, i.e.

$$\hat{u}_t = (1-L)^{d_0} y_t - \hat{\beta}' w_t; \quad \hat{\beta} = \left( \sum_{t=1}^T w_t w_t' \right)^{-1} \sum_{t=1}^T w_t (1-L)^{d_0} y_t; \quad w_t = (1-L)^{d_0} z_t;$$

where the function  $g$  is a known function coming from the spectral density function of  $u_t$ ,

$$f(\lambda; \sigma^2; \tau) = \frac{\sigma^2}{2\pi} g(\lambda; \tau), \quad -\pi < \lambda \leq \pi.$$

Note that these tests are of a purely parametric nature and therefore require specific modeling assumptions regarding the short memory specification of  $u_t$ . Thus, for example, if  $u_t$  is white noise,  $g \equiv 1$ , and if  $u_t$  is an AR process of the form  $\phi(L)u_t = \varepsilon_t$ ,  $g = |\phi(e^{i\lambda})|^{-2}$ , with  $\sigma^2 = V(\varepsilon)$ , so that the AR coefficients are a function of  $\tau$ .

Based on the null hypothesis  $H_0$  (2), Robinson (1994) showed that under certain regularity conditions:<sup>1</sup>

$$\hat{r} \rightarrow_d N(0, 1) \quad \text{as } T \rightarrow \infty. \quad (5)$$

He also proves the Pitman efficiency theory of the tests against local departures from the null. Thus, we are in a classical large sample-testing situation: an approximate one-sided  $100\alpha\%$ -level test of  $H_0$  (2) against the alternative:  $H_a: d > d_0$  ( $d < d_0$ ) will be given by the rule: "Reject  $H_0$  if  $\hat{r} > z_\alpha$  ( $\hat{r} < -z_\alpha$ )," where the probability that a standard normal variate exceeds  $z_\alpha$  is  $\alpha$ . This version of Robinson's (1994) tests was used in empirical applications in Gil-Alana and Robinson (1997) and Gil-Alana (2000). Other versions of Robinson's tests, based on seasonal (quarterly and monthly), and cyclical data can be found in Gil-Alana and Robinson (2001) and Gil-Alana (1999, 2001a) respectively.

#### 4. Testing the Order of Integration in the Short-run Interest Rates

The time-series data for the monthly seasonally unadjusted short-run interest rates,<sup>2</sup> for Mexico, Korea, Malaysia, the Philippines, Singapore, and Thailand, are extracted from the International Financial Series (IFS) of the International Monetary Fund (IMF). The series start in January 1980 in all countries and end in September 2001 (Korea and Malaysia), October 2001 (Singapore), November 2001 (Philippines and Thailand), and December 2001 (Mexico).

Figure 1 displays plots of the original time series. The common characteristic among the countries is that they all have suffered from currency crises (the Asian crisis in 1997 and the Mexican tequila crisis in 1994). Nevertheless, they differ in terms of regime of change (Mexico and Malaysia have pegged their exchange rate to the US dollar during long periods, whereas the exchange rate has been less controlled during this period in the other countries)<sup>3</sup> and in terms of stabilization policy: when the exchange rate is fixed, a change in the interest rate is the only monetary instrument to stabilize the

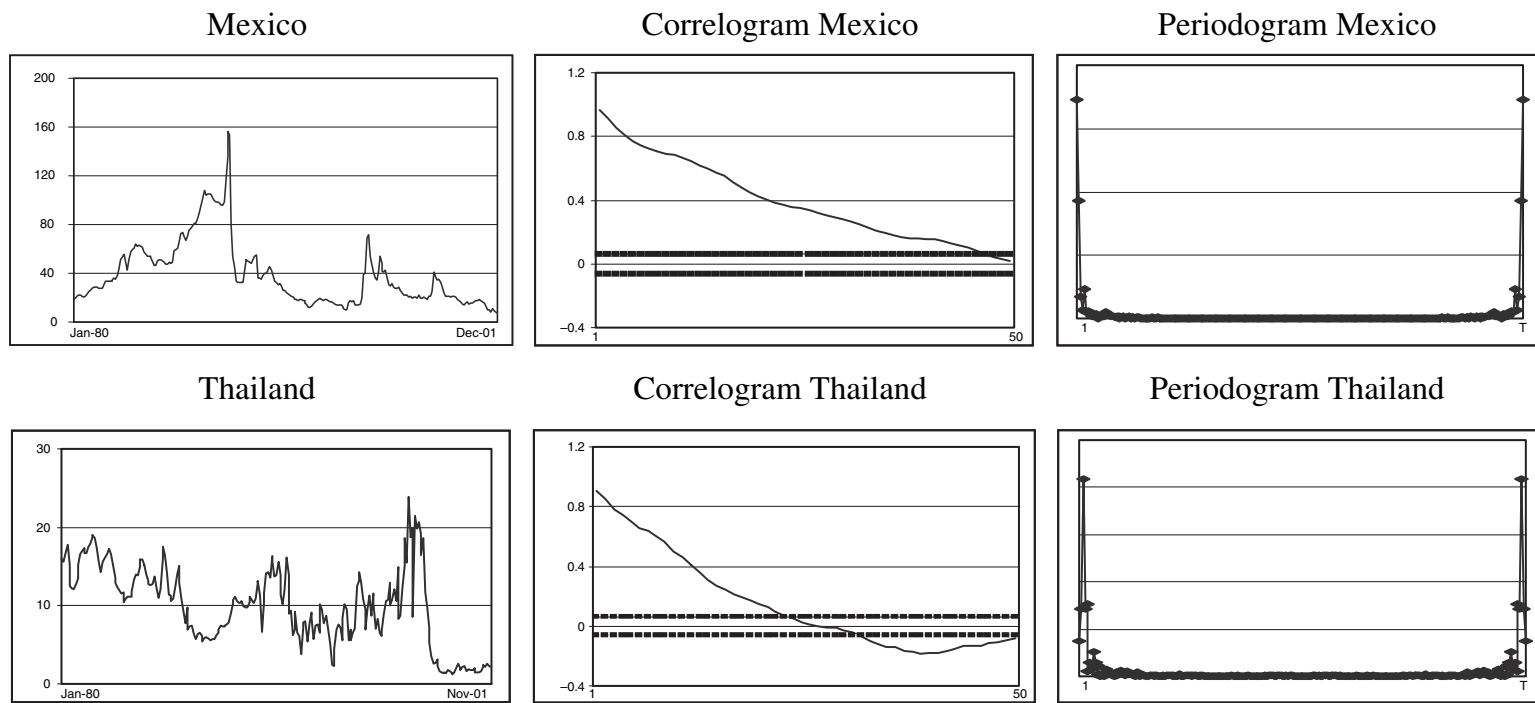


Figure 1. Plots of the Original Series with their Corresponding Correlograms and Periodograms

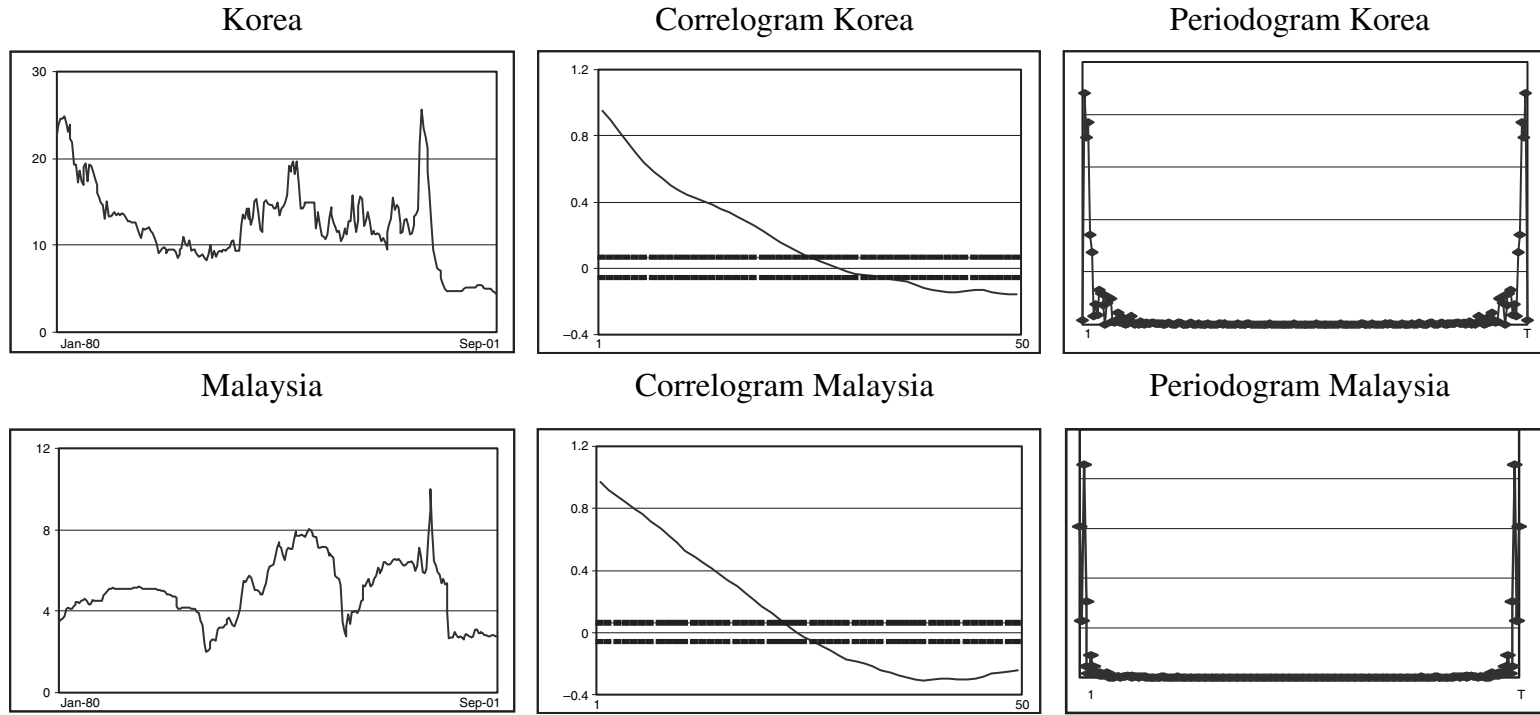
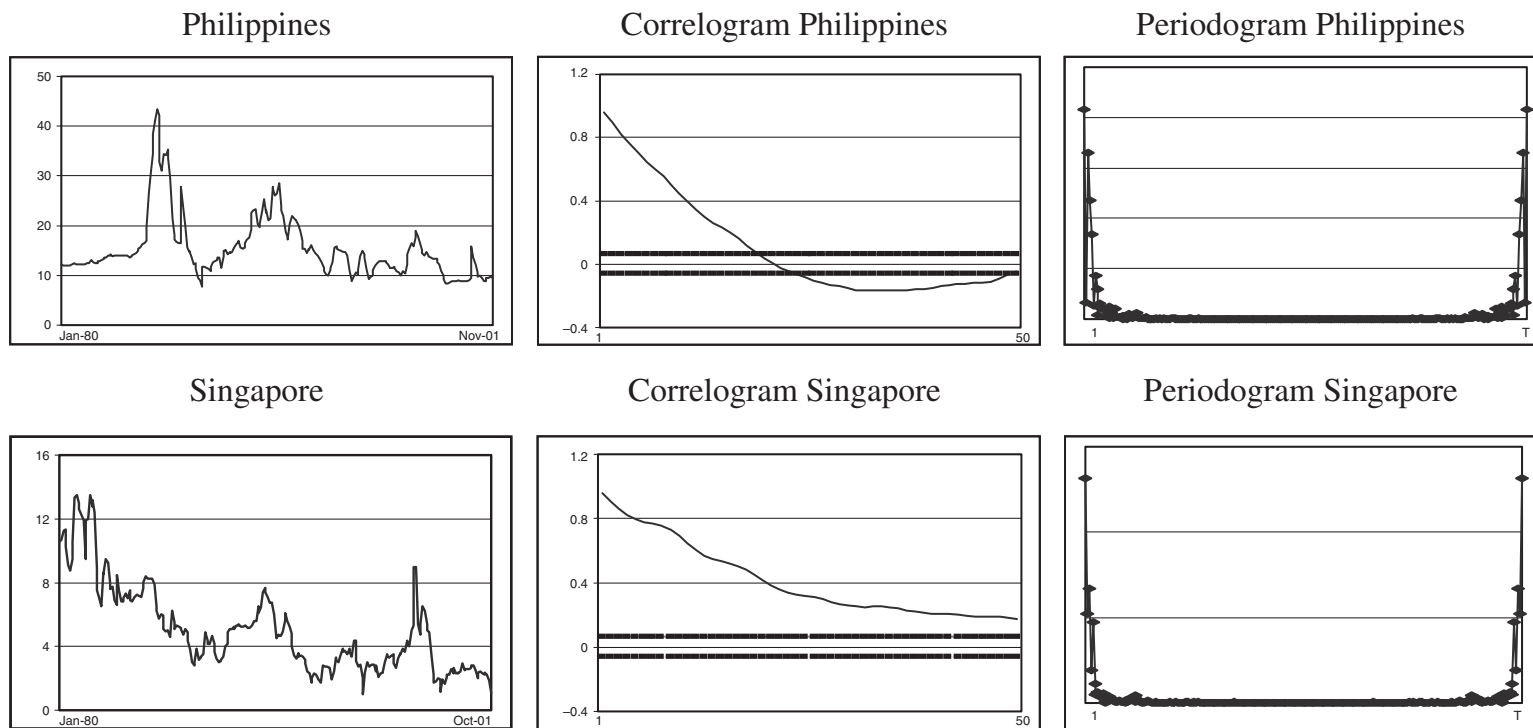


Figure 1. Continued



Note: The large-sample standard error under the null hypothesis of no autocorrelation is  $1/\sqrt{T}$ , or roughly 0.061.

Figure 1. Continued

economy after a currency crisis. This property clearly appears in Figure 1 for Mexico in 1994 during the crisis. As the Mexican peso was pegged to the US dollar, Mexican authorities were forced to increase the short-run interest rates to 160% per annum to avoid a complete outflow of capital. The succession of periods of complete peg, controlled peg, or perfect float could lead to the appearance of structural breaks. In such a case, the persistence of the series will be overvalued by traditional tests (see Diebold and Inoue, 2001). But as far as we know, there are no fractional integration tests that are robust to structural breaks at an unknown date and so we leave this topic for future research.

In Figure 1, we observe that all interest rates are nonstationary. This feature can also be visualized in the correlograms (with autocorrelations decreasing very slowly) and in the periodograms (with very large values at the smallest frequencies). Plots of the first differences of the series, with their corresponding correlograms and periodograms, are displayed in Figure 2. The series may now be stationary, though the correlograms still show significant values even at lags relatively far away from zero, which might signal that fractional differencing with a value of  $d$  smaller than or greater than 1 may be more appropriate than first differences. In addition, the periodograms in some of the countries (Thailand, Korea, the Philippines, and Singapore) show values close to 0 at the zero frequency, which might suggest that these series are now overdifferenced.

Denoting each of the time series by  $y_t$ , we use the model given by (1) and (3), with  $z_t = (1, t)'$ ,  $t \geq 1$ ,  $z_t = (0, 0)'$  else. Thus, under the null hypothesis  $H_0$  (2):

$$y_t = \beta_0 + \beta_1 t + x_t, \quad t = 1, 2, \dots \quad (6)$$

$$(1 - L)^{d_0} x_t = u_t, \quad t = 1, 2, \dots \quad (7)$$

We treat separately the cases  $\beta_0 = \beta_1 = 0$  *a priori*;  $\beta_0$  unknown and  $\beta_1 = 0$  *a priori*; and  $\beta_0$  and  $\beta_1$  unknown, i.e. we consider respectively the cases of no regressors in the undifferenced regression (6), an intercept, and an intercept and a linear time trend, and report the test statistic, not merely for the case of  $d_0 = 1$  (a unit root), but for  $d_0 = 0.50$  (0.10), 1.50, thus also including a test for stationarity ( $d_0 = 0.5$ ) as well as other fractionally integrated possibilities.

The test statistic reported in Table 1 (and also in Tables 2 and 3) is the one-sided statistic corresponding to  $\hat{r}$  in (4), so that significantly positive values are consistent with orders of integration higher than  $d_0$ , whereas significantly negative values are consistent with alternatives:  $d < d_0$ . A notable feature observed in Table 1, in which  $u_t$  is taken to be white noise, is that the values of the test statistic monotonically decrease with  $d_0$ . This is to be expected because they are one-sided statistics. Thus, for example, if  $H_0$  (2) is rejected with  $d_0 = 1$  against alternatives of form:  $H_a: d > 1$ , an even more statistically significant outcome in this direction should be expected for testing the null hypothesis  $d_0 = 0.75$  or  $d_0 = 0.50$ . Starting with the “no regressors” case in Table 1(i), we see that the unit-root null hypothesis is rejected for Mexico, Malaysia, the Philippines, and Korea in favor of alternatives with  $d$  higher than 1. For Korea and Singapore, the unit root cannot be rejected, though in the former country,  $d_0 = 1.10$  is also not rejected whilst  $d_0 = 0.90$  is not rejected for Singapore. Finally, the series for Thailand seems to be the closest to stationary, because  $H_0$  (2) cannot be rejected when  $d_0 = 0.80$  and  $0.90$ . Tables 1(ii) and (iii) report results with, respectively,  $\beta_1 = 0$  *a priori* (no time trend in the undifferenced regression), and both  $\beta_0$  and  $\beta_1$  unrestricted. In all cases,  $\hat{r}$  is monotonic and, moreover, while there are some differences in the values of  $\hat{r}$  across Tables 1(ii) and (iii) for the same series/ $d_0$  combination, the conclusions suggested by both seem very similar with the nonrejection values occurring at the same



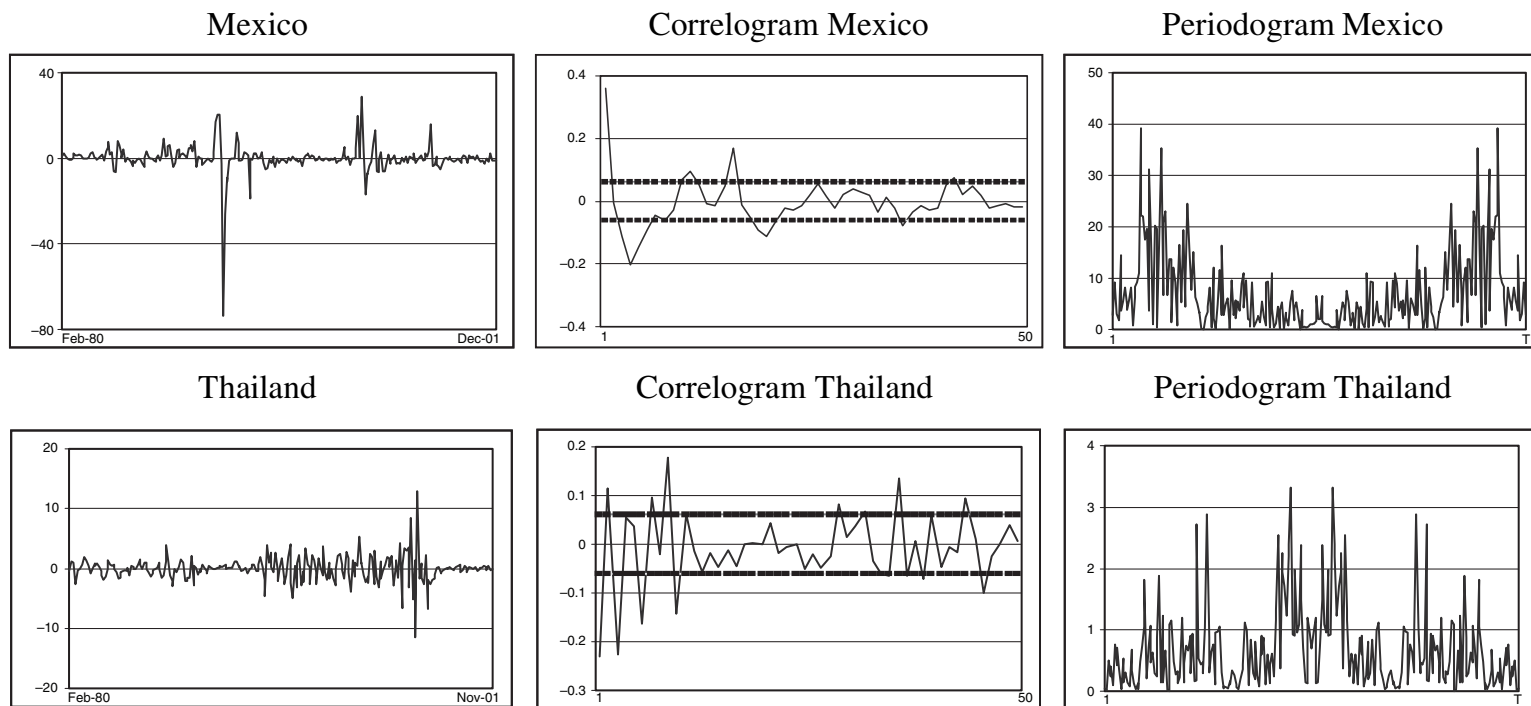


Figure 2. Plots of the First-differenced Series, with their Corresponding Correlograms and Periodograms

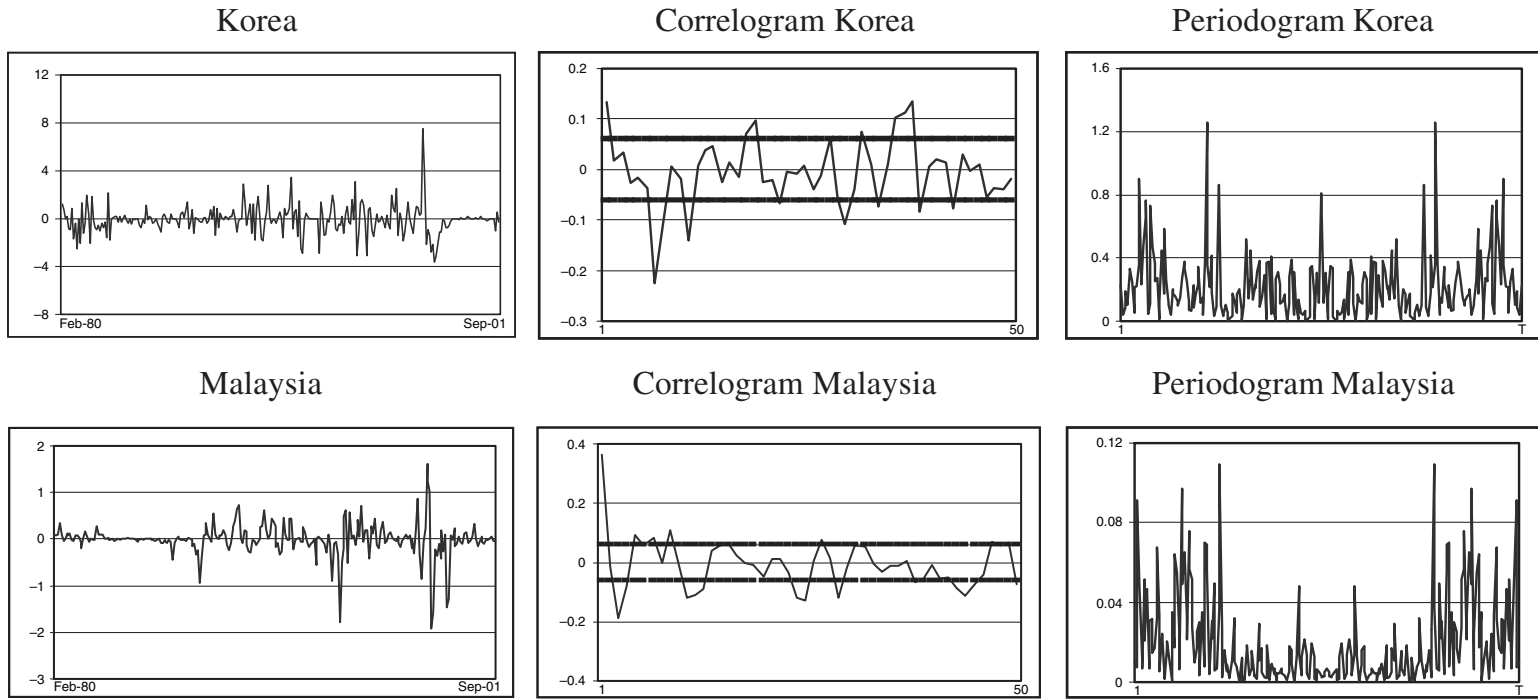
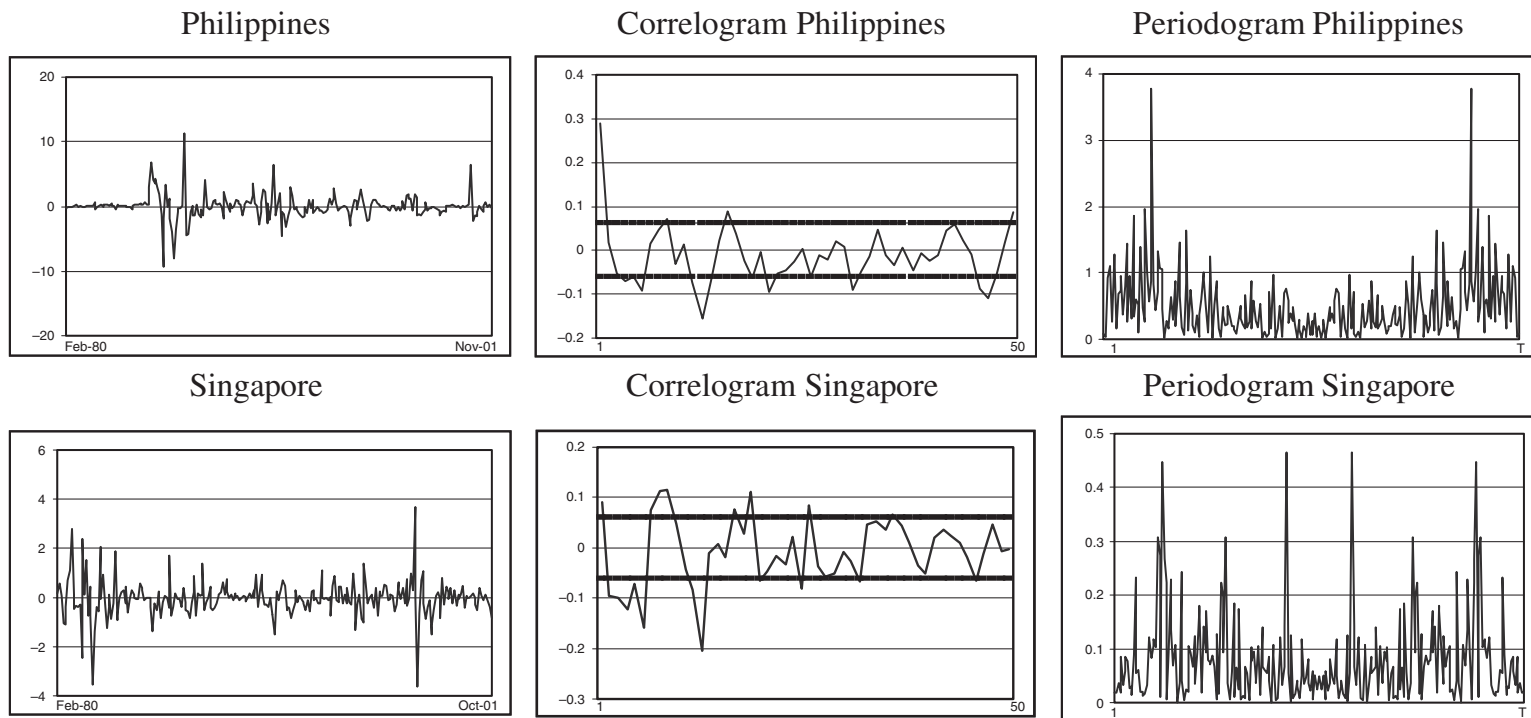


Figure 2. *Continued*



Note: The large-sample standard error under the null hypothesis of no autocorrelation is  $1/\sqrt{T}$ , or roughly 0.061.

Figure 2. Continued

Table 1. Testing  $H_0$  (2) in (1) and (3) with White Noise Disturbances

		(i) With no regressors									
Country/ $d$	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	18.43	13.76	9.91	6.92	4.63	2.87	<b>1.47</b>	<b>0.32</b>	<b>-0.64</b>	<b>-1.48</b>	-2.21
Malaysia	19.96	16.03	12.07	8.43	5.37	2.92	<b>1.03</b>	<b>-0.42</b>	<b>-1.56</b>	-2.46	-3.19
Philippines	17.74	13.83	10.23	7.12	4.53	2.42	<b>0.72</b>	<b>-0.64</b>	-1.75	-2.66	-3.41
Korea	14.58	11.04	7.75	4.89	2.51	<b>0.57</b>	<b>-0.98</b>	-2.23	-3.24	-4.04	-4.70
Singapore	14.86	10.20	6.27	3.20	<b>0.88</b>	<b>-0.83</b>	-2.12	-3.11	-3.87	-4.49	-4.99
Thailand	10.88	6.68	3.24	<b>0.59</b>	<b>-1.37</b>	-2.82	-3.91	-4.73	-5.37	-5.87	-6.28
		(ii) With an intercept									
Country/ $d$	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	17.24	13.05	9.57	6.81	4.65	2.95	<b>1.58</b>	<b>0.45</b>	<b>-0.51</b>	<b>-1.35</b>	-2.08
Malaysia	20.79	16.62	12.66	9.14	6.19	3.84	2.00	<b>-0.57</b>	<b>-0.56</b>	<b>-1.47</b>	-2.21
Philippines	16.65	13.26	10.12	7.33	4.94	2.94	<b>1.27</b>	<b>-0.09</b>	<b>-1.22</b>	-2.15	-2.93
Korea	14.39	11.38	8.37	5.55	3.13	<b>1.16</b>	<b>-0.42</b>	-1.70	-2.74	-3.58	-4.28
Singapore	10.33	6.66	4.14	2.20	<b>0.63</b>	<b>-0.64</b>	-1.68	-2.54	-3.26	-3.87	-4.39
Thailand	8.05	4.42	<b>1.61</b>	<b>-0.53</b>	-2.16	-3.39	-4.33	-5.05	-5.62	-6.08	-6.45
		(iii) With an intercept and a linear time trend									
Country/ $d$	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	16.68	12.83	9.49	6.78	4.64	2.95	<b>1.58</b>	<b>0.44</b>	<b>-0.51</b>	<b>-1.35</b>	-2.08
Malaysia	20.80	16.62	12.65	9.12	6.19	3.84	2.00	<b>-0.56</b>	<b>-0.56</b>	<b>-1.47</b>	-2.22
Philippines	16.42	13.16	10.08	7.32	4.94	2.94	<b>1.28</b>	<b>-0.08</b>	<b>-1.22</b>	-2.15	-2.93
Korea	16.66	12.66	8.91	5.73	3.18	<b>1.16</b>	<b>-0.43</b>	-1.71	-2.74	-3.59	-4.27
Singapore	10.51	7.13	4.43	2.32	<b>0.66</b>	<b>-0.63</b>	-1.69	-2.55	-3.26	-3.87	-4.39
Thailand	7.76	4.39	<b>1.59</b>	<b>-0.54</b>	-2.16	-3.39	-4.33	-5.05	-5.62	-6.08	-6.45

Note: The non-rejection values of the null hypothesis at the 5% significance level are shown in bold.

values of  $d_0$  in both tables. The results are also similar to those given in Table 1(i): Mexico appears to be the most nonstationary series, with the nonrejection values of  $d$  ranging between 1.10 and 1.40, followed by Malaysia and the Philippines, with values of  $d$  higher than 1. The unit root cannot be rejected for Korea and Singapore, while Thailand is the closest to stationarity series, with nonrejection for  $d_0 = 0.7$  and 0.8.

However, the significance of these results may be due to a large extent to the zero autocorrelation assumption in  $u_t$ . Therefore, we also performed the tests imposing autoregressive disturbances. The results for  $u_t$  following an AR(1) are listed in Table 2. If we do not include regressors (Table 2(i)), there is a lack of monotonicity in the value of the test statistic with respect to  $d_0$  for all series, except for Mexico. This lack of monotonicity may be an indication of model misspecification (see e.g. Gil-Alana and Robinson, 1997), but this may also be due to the fact that the AR coefficients are Yule–Walker estimates. In this case, although the coefficients are smaller than 1 in absolute value, they can be arbitrarily close to 1. Thus, a problem may then occur because they could capture the order of integration of the series by means, for example, of a coefficient of 0.99 when using AR(1) disturbances. When including an intercept, monotonicity is achieved for Mexico and the Philippines and, while including a linear time trend, this property is satisfied for all series except Malaysia and the Philippines. In the case of AR(1) disturbances and a linear time trend, the unit-root null hypothesis cannot be rejected for Thailand. However, most of the nonrejections take place when  $d$  is smaller than 1, suggesting that if the disturbances are autocorrelated, the

Table 2. Testing  $H_0$  (2) in (1) and (3) with AR(1) Disturbances

(i) With no regressors											
Country/d	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	2.17	<b>1.29</b>	<b>-0.35</b>	-1.77	-2.66	-3.19	-3.52	-3.73	-3.88	-4.01	-4.13
Malaysia	-3.96	-5.39	-5.87	-4.67	-3.61	-3.19	-3.14	-3.25	-3.43	-3.62	-3.81
Philippines	-2.79	-3.59	-3.36	-2.79	-2.51	-2.48	-2.61	-2.83	-3.08	-3.35	-3.62
Korea	-10.43	-14.40	-13.65	-8.30	-4.51	-2.46	-1.51	-1.25	-1.40	-1.77	-2.23
Singapore	-11.13	-11.72	-7.51	-4.98	-3.78	-3.12	-3.01	-3.04	-3.19	-3.39	-3.62
Thailand	-0.63	1.04	1.28	0.86	0.18	-0.55	-1.24	-1.86	-2.41	-2.86	-3.32
(ii) With an intercept											
Country/d	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	3.12	2.16	<b>0.28</b>	<b>-1.32</b>	-2.35	-2.99	-3.39	-3.66	-3.85	-4.01	-4.16
Malaysia	0.05	1.32	1.44	0.37	-0.85	-1.87	-2.68	-3.29	-3.76	-4.10	-4.36
Philippines	<b>-0.35</b>	<b>-0.42</b>	<b>-0.76</b>	<b>-1.17</b>	<b>-1.58</b>	-2.00	-2.41	-2.80	-3.16	-3.49	-3.79
Korea	-6.99	-3.03	-0.78	-0.47	-0.74	-1.18	-1.65	-2.14	-2.62	-3.09	-3.59
Singapore	-4.00	-1.93	-1.80	-2.12	-2.52	-2.93	-3.28	-3.59	-3.87	-4.12	-4.35
Thailand	1.67	1.82	1.17	0.33	-0.49	-1.25	-1.92	-2.48	-2.97	-3.38	-3.74
(iii) With an intercept and a linear time trend											
Country/d	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	2.90	1.83	<b>0.11</b>	<b>-1.37</b>	-2.36	-2.99	-3.39	-3.66	-3.86	-4.02	-4.16
Malaysia	1.78	1.97	1.47	0.31	-0.87	-1.87	-2.67	-3.29	-3.76	-4.11	-4.37
Philippines	0.05	<b>-0.47</b>	<b>-0.87</b>	<b>-1.22</b>	-1.60	-2.00	-2.41	-2.80	-3.16	-3.49	-3.79
Korea	-1.10	0.15	0.27	-0.14	-0.66	-1.17	-1.67	-2.16	-2.63	-3.09	-3.52
Singapore	<b>0.57</b>	<b>-0.24</b>	<b>-1.16</b>	-1.90	-2.47	-2.92	-3.29	-3.60	-3.87	-4.12	-4.35
Thailand	2.29	1.93	<b>1.19</b>	<b>0.33</b>	<b>-0.49</b>	<b>-1.25</b>	-1.92	-2.48	-2.97	-3.38	-3.74

Note: The non-rejection values of the null hypothesis at the 5% significance level are shown in bold.

order of integration of the series is smaller. It therefore indicates that the AR model is somewhat confounded with the fractional one in finite samples. Using MA disturbances, the results did not substantially differ from those given in Table 1 for the case of white noise. Higher-order AR (and ARMA) processes were also tested and the lack of monotonicity became even more apparent as we increase the order of the AR process.

In order to solve this problem, we used another type of  $I(0)$  disturbances as proposed by Bloomfield (1973) and which accommodates fairly well to the present version of the tests. Using this method, the disturbances are exclusively specified in terms of the spectral density function:

$$f(\lambda; \tau) = \frac{\sigma^2}{2\pi} \exp\left(2 \sum_{r=1}^m \tau_r \cos(\lambda r)\right). \quad (8)$$

Bloomfield (1973) showed that the logarithm of the spectral density function for an  $\text{ARMA}(p, q)$  process can be approximated by a truncated Fourier series. He showed that (8) approximates the spectrum of an ARMA process well where  $p$  and  $q$  are small values, which is usually the case in practice. Like the stationary  $\text{AR}(p)$  case, this model has exponentially decreasing autocorrelations and thus we do not need to rely on as many parameters as in the ARMA processes, which always turn out to be tedious in terms of estimation, testing, and model specification. Furthermore, unlike AR processes,  $u_t$  is stationary for all real values of  $\tau$ . Formulas for the Newton-type

Table 3. Testing  $H_0(2)$  in (1) and (3) with Bloomfield (1) Disturbances

(i) With no regressors											
Country/ <i>d</i>	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	5.09	2.77	<b>0.73</b>	<b>-0.57</b>	<b>-1.57</b>	-2.38	-2.91	-3.27	-3.66	-3.82	-4.05
Malaysia	6.72	5.00	3.22	<b>1.55</b>	<b>0.09</b>	<b>-1.14</b>	-2.00	-2.76	-3.28	-3.79	-4.16
Philippines	5.57	3.71	2.32	<b>0.82</b>	<b>-0.24</b>	<b>-1.19</b>	-1.86	-2.51	-2.99	-3.35	-3.75
Korea	5.28	3.50	2.39	<b>1.07</b>	<b>0.10</b>	<b>-0.81</b>	<b>-1.49</b>	-1.98	-2.52	-2.94	-3.27
Singapore	5.61	3.64	1.66	<b>0.26</b>	<b>-0.87</b>	-1.81	-2.61	-3.15	-3.51	-3.89	-4.15
Thailand	5.83	3.94	2.23	<b>1.01</b>	<b>-0.27</b>	<b>-1.18</b>	-1.80	-2.58	-2.99	-3.44	-3.77
(ii) With an intercept											
Country/ <i>d</i>	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	4.19	2.00	<b>0.53</b>	<b>-0.84</b>	-1.74	-2.31	-2.94	-3.33	-3.57	-3.87	-4.10
Malaysia	6.67	4.40	2.37	<b>0.87</b>	<b>-0.41</b>	<b>-1.49</b>	-2.40	-3.03	-3.60	-4.01	-4.30
Philippines	4.64	2.85	<b>1.52</b>	<b>0.30</b>	<b>-0.57</b>	<b>-1.39</b>	-2.00	-2.60	-3.05	-3.54	-3.77
Korea	4.50	3.20	2.03	<b>0.77</b>	<b>-0.18</b>	<b>-0.86</b>	<b>-1.52</b>	-2.18	-2.69	-3.09	-3.40
Singapore	2.23	<b>0.29</b>	<b>-0.69</b>	-1.68	-2.23	-2.79	-3.19	-3.62	-3.96	-4.21	-4.40
Thailand	4.40	2.89	<b>1.37</b>	<b>0.29</b>	<b>-0.68</b>	<b>-1.59</b>	-2.27	-2.76	-3.27	-3.65	-3.95
(iii) With an intercept and a linear time trend											
Country/ <i>d</i>	0.50	0.60	0.70	0.80	0.90	1.00	1.10	1.20	1.30	1.40	1.50
Mexico	3.93	1.79	<b>0.45</b>	<b>-0.86</b>	-1.74	-2.31	-2.94	-3.34	-3.57	-3.87	-4.10
Malaysia	6.70	4.41	2.36	<b>0.86</b>	<b>-0.41</b>	<b>-1.44</b>	-2.40	-3.03	-3.60	-4.02	-4.31
Philippines	4.41	2.76	<b>1.48</b>	<b>0.29</b>	<b>-0.57</b>	<b>-1.39</b>	-2.00	-2.60	-3.05	-3.54	-3.77
Korea	5.88	4.05	2.55	<b>0.96</b>	<b>-0.13</b>	<b>-0.86</b>	<b>-1.54</b>	-2.20	-2.70	-3.09	-3.38
Singapore	2.20	<b>0.80</b>	<b>-0.59</b>	<b>-1.54</b>	-2.19	-2.78	-3.20	-3.64	-3.97	-4.21	-4.40
Thailand	4.31	2.77	<b>1.34</b>	<b>0.29</b>	<b>-0.68</b>	<b>-1.59</b>	-2.27	-2.76	-3.27	-3.66	-3.95

Note: The non-rejection values of the null hypothesis at the 5% significance level are shown in bold.

iteration for estimating the  $\tau_r$  are very simple (involving no matrix inversion). Updating formulas when  $m$  is increased are also simple, and we can replace  $\hat{A}$  below (4) by the population quantity:

$$\sum_{l=m+1}^{\infty} l^{-2} = \frac{\pi^2}{6} - \sum_{l=1}^m l^{-2},$$

which is indeed constant with respect to the  $\tau_r$  (unlike in the AR case). The Bloomfield model for  $I(0)$  processes, combined with the fractional model (1) has not been used extensively in previous econometric applications (though the Bloomfield model itself is a well-known model in other disciplines, e.g. Beran, 1993). One byproduct of this work is its emergence as a credible alternative to the fractional ARIMAs, which have become conventional in parametric modeling of long memory (see Gil-Alana and Robinson, 1997; Velasco and Robinson, 1999; Gil-Alana, 2001b).

Table 3 shows the results based on Bloomfield (1973) disturbances ( $m = 1$ ). We also tried other values for  $m$  and the results were very similar to those reported in Table 3. Monotonicity is now achieved for all series, independent of the inclusion of deterministic trends. We see here that the values of  $d_0$  where  $H_0(2)$  cannot be rejected, range between 0.6 and 1.1. The fact that  $H_0(2)$  is rejected with  $d = 0.5$  in favor of alternatives of form:  $d > 0.5$  also suggests that the series are nonstationary and clearly reject the trend-stationary alternatives suggested by some authors.

Table 4 summarizes the results from Tables 1 and 3 by means of reporting the 95% confidence intervals of the values of  $d_0$  where  $H_0(2)$  cannot be rejected along with the

Table 4. Confidence Intervals and Values of  $d_0$  which Produces the Lowest Statistics Across  $d_0$ 

Country	$u_t$	No regressor		An intercept		A linear time trend	
		Confidence interval	$d_0$	Confidence interval	$d_0$	Confidence interval	$d_0$
Mexico	White noise	[1.09–1.42]	1.23	[1.10–1.43]	1.24	[1.10–1.43]	1.24
	Bloomfield	[0.65–0.91]	0.75	[0.62–0.88]	0.74	[0.62–0.88]	0.74
Malaysia	White noise	[1.07–1.31]	1.17	[1.13–1.42]	1.25	[1.13–1.42]	1.25
	Bloomfield	[0.78–1.06]	0.91	[0.76–1.00]	0.86	[0.76–1.00]	0.86
Philippines	White noise	[1.05–1.28]	1.15	[1.08–1.34]	1.19	[1.08–1.34]	1.19
	Bloomfield	[0.74–1.04]	0.87	[0.70–1.03]	0.83	[0.71–1.03]	0.83
Korea	White noise	[0.94–1.15]	1.03	[0.98–1.19]	1.07	[0.98–1.19]	1.07
	Bloomfield	[0.76–1.14]	0.93	[0.73–1.10]	0.90	[0.76–1.10]	0.91
Singapore	White noise	[0.87–1.05]	0.95	[0.84–1.09]	0.95	[0.84–1.09]	0.95
	Bloomfield	[0.71–0.98]	0.82	[0.53–0.79]	0.62	[0.54–0.81]	0.64
Thailand	White noise	[0.76–0.91]	0.83	[0.70–0.86]	0.77	[0.70–0.86]	0.77
	Bloomfield	[0.76–1.05]	0.88	[0.67–1.01]	0.82	[0.67–1.01]	0.82

values of  $d_0$ , which produces the lowest statistics in absolute value across  $d_0$ . The first thing we observe is that the values are smaller if the disturbances are autocorrelated, suggesting that some competition might exist between the fractional differencing parameter and the autocorrelated disturbances in describing the nonstationary component of the series. According to the summary of results for the individual series we conclude that the interest rates in Mexico and Malaysia exhibit the higher degree of nonstationarity, followed by the Philippines and Korea. In all these cases, the orders of integration are higher than 1 if the disturbances are white noise but smaller than 1 if they are autocorrelated. On the other hand, the series corresponding to Singapore and Thailand are clearly mean-reverting, with orders of integration smaller than 1 in both the cases of white noise and autocorrelated disturbances.

## 5. Concluding Comments

In this article we have examined the stochastic behavior of the short-run interest rates in six emerging countries by means of fractional integration techniques. More specifically we used a version of Robinson's (1994) test that allows for testing  $I(d)$  statistical models. These tests have several distinguishing features compared to other procedures for testing unit and/or fractional roots. In particular, they have a standard null limiting distribution, that is also unaffected by the inclusion of deterministic trends and different types of  $I(0)$  disturbances. In addition, the tests are the most efficient ones, for the appropriate (fractional) alternatives. The results show that the interest rates in Mexico, Malaysia, the Philippines, and Korea present orders of integration higher than 1 if the disturbances are white noise and smaller than 1 if they are uncorrelated. The orders of integration in Singapore and Thailand are smaller than 1 in all cases, suggesting that they are mean-reverting. Thus UIP seems to hold and the short-run interest rates have been kept in line with the constraints imposed by the capital market

even during the Asian 'flu scare. The results are less clear for Mexico, Malaysia, the Philippines, and Korea, where interest rates are often found to be nonstationary. The latter result sheds some light on the stabilization policy after the currency crises: whereas Singapore and Thailand used stabilization policies based on exchange rates (which exhibit a higher volatility: see Candelon and Straetmans, 2003), the other countries performed more active restrictive monetary policies leading to long memory in the interest rates.

Several other lines of research are in progress, which should prove relevant to the analysis of these and other macroeconomic time series. Multivariate versions of the tests of Robinson (1994) are being developed and this can lead to an alternative approach to the study of cointegration. The Bloomfield (1973) model for the  $I(0)$  components is also currently being developed in a multivariate set-up.

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## Notes

1. These conditions are very mild regarding technical assumptions, which are satisfied by models (1) and (3).
2. Short-run interest rates correspond to the call rates.
3. A complete chronology of economic, political, and financial events in these countries can be found in the homepage of C. R. Harvey at [http://duke.edu/~charvey/country\\_risk/chronology/](http://duke.edu/~charvey/country_risk/chronology/).